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Juan J Dolado and Ramón María-Dolores

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Juan J Dolado, Universidad Carlos III de Madrid and CEPR
Ramón María-Dolores, Universidad de Murcia

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Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: http://www.cepr.org

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ABSTRACT

Evaluating Changes In The Bank Of Spain's Intervention: An Alternative Approach Using Marked Point Processes*

In this Paper we provide empirical evidence on the determinants of the monetary policy stance by the Bank of Spain over the period 1984–98, by means of modelling a marked point process explaining the probability of an intervention at each point in time (events) and the size of these interventions (marks) conditional on the decision to intervene. Interventions are measured by changes in the marginal interest rate of the Spanish daily interbank market. We test for and find evidence in favour of the existence of asymmetries in the response of the central bank to the evolution of various macroeconomic variables and for the presence of 'duration' effects.

JEL Classification: E52, E58
Keywords: marked point process, probit , ordered probit, sequential probit, monetary policy

Juan J Dolado
Dpto. de Economía
Universidad Carlos III de Madrid
c/ Madrid, 126 28903
Getafe (Madrid)
SPAIN
Tel: (349 16) 24 93 00
Fax: (349 16) 24 93 13
Email: dolado@eco.uc3m.es

Ramón María-Dolores
Dpto. de Fundamentos del Análisis Económico
Universidad de Murcia
Facultad de Economía y Empresa
Campus de Espinardo 30300
Murcia
SPAIN
Tel: (349 16) 24 97 54
Fax: (349 16) 24 98 75
Email: ramonmar@eco.uc3m.es

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NON-TECHNICAL SUMMARY

The goal of this Paper is to characterize the evolution of the Bank of Spain's interventions in the daily interbank market through changes in a target marginal interest rate, using a methodology based on the analysis of the so-called *marked point processes*. In the statistics literature, a process that evolves stochastically in time and has associated to it an information process at each event time is known as a market point process. The strategy of the Paper is to break down the problem in two parts. First, we model the timing of interventions (*events*) and then model the changes in the interest rate (*marks*), all conditional on past information.

Since the interest rate is changed irregularly in time and changes take place in discrete increments, the use of such a methodology allows us to address some interesting issues that range from *What determines the arrival times and the size of interventions?* to *Are there asymmetric responses of the Bank of Spain to the evolution of the determinants of interventions?*

To study the determinants of events we estimate a \QTR{em}{probit} model determining the probability of intervening or not intervening as a function of various macroeconomic indicators that the central bank observes when making the decision and the time elapsed from the last intervention. As for marks, an *ordered probit* is estimated to determine the size of the target change, conditional on having decided to intervene. Finally, a *sequential probit* model is estimated to study the joint decision of intervening and the corresponding target change. The latter approach serves as a benchmark to test for the consistency of the estimate obtained through the separate procedure. It should be noticed that although the application in this Paper is confined to the analysis of the behaviour of a single central bank, the methodological scope of the Paper can be easily extended to more general cases.

The sample period is 1984–98 and covers the period when the Bank of Spain used an interest rate as its main instrument of monetary policy before handing over its control to the European Central Bank.

Over that period there have been 700 interventions. The average duration between interventions is about 8 days although there is an outlier of 181 days when the transition from a monetary aggregate-based control scheme to an interest rate rule was being completed. As for the determinants we chose the following macroeconomic variables at monthly frequency: changes of the inflation rate, changes of the pta/DM exchange rate, deviations of money growth over its target rate, changes in the Industrial Production Index, changes in lagged interest rates and the time elapsed from the last intervention.

The results regarding the events point out that there is a higher probability of intervention when: (i) there is a recession, (ii) inflation accelerates, (iii) the exchange rate depreciates, (iv) when the interest rate increased in previous periods and (v) when there has been a long period without interventions. In particular, the effect of inflation on such a decision is much larger since 1992.

As regards the results on marks, we find evidence in favour of various types of ‘duration’ effects. First, longer inactivity periods induce stronger changes in interest rates than frequent adjustments. Secondly, the evidence points out that the Bank of Spain was more prone to avoid overheating, by implementing large increases in interest rates during expansions than to avoid recessions through large cuts in interest rates.

1 Introduction

Traditionally, as in many other countries (see Clarida et al, 1998), the analysis of the determinants of the Spanish monetary policy stance has mostly focused on the estimation of reaction functions of the Bank of Spain through a Taylor(1993) rule specification where interventions in terms of changes of a short interest rate are modelled as function of a set of explanatory variables; c.f. María-Dolores(1998) and Galí(1999).

In this paper we complement the previous approach by characterising the evolution of the Bank of Spain´s interventions using an alternative methodology based on the analysis of the so-called *marked point processes*(see Engle and Russell, 1995, 1997) . According to these authors a *marked point process* is a sequence of orderly times where a number of events occur at each point in time and where there are characteristics associated with the arrival times such as a price or volume change.

The fact that: (i) changes in interest rate do not occur regularly in time and (ii) changes take place in discrete increments, favours the use of such a methodological approach to address a number of interesting issues regarding the evolution of the monetary policy stance. Those issues range from *What determines the arrival times and the size of interventions?* to *Are there asymmetric responses of the Bank of Spain to the evolution of the determinants of interventions?*, which cannot be dealt with the standard Taylor(1993) rule´s approach. To address those issues, we will estimate a probit model to determine the probability of an intervention at each point in time (the so-called *events*) and an ordered probit to determine the size of the intervention, conditional on having decided to intervene (the so-called *marks*). Separate and joint estimates of the parameters involving both decisions will be compared in order to assess their robustness to the choice of estimation method.

To our knowledge, the only empirical studies which address similar questions using this type of econometric techniques are Jordá(1998) and Fischer and Zurlinden(1999). However, on the one hand, the latter authors only study the determinants of *events* at the Federal Reserve, the Bundesbank and the Swiss National Bank by means of an Autorregresive Conditional Duration (ACD) model, ignoring therefore the analysis of *marks*, whereas, on the other hand, the former author models the duration time between *events*(again through an ACD model) and the evolution of *marks* in a somewhat similar manner to the one used in this paper, but without dealing with their joint modellisation. Moreover, none of the two papers deal with the issue of asymmetries in central bank´s intervention. In view of the above, our paper´s contribution to this literature may be interesting for two reasons: (i) by comparing the estimates obtained through separate and joint modelling of both decisions, a check on the consistency of the estimates obtained under the first method is provided, and (ii) by testing for the existence of asymmetries in central bank´s interventions, new evidence is obtained on a hypothesis of growing interest in recent empirical and theoretical work on the effects of monetary policy(see Ravn and Sola, 1996). Moreover, through the application is restricted to analyse the behaviour of a single central bank, the methodological scope of the paper could be easily extended to more general cases.

Proceeding in that way, we obtain several interesting results. First, there is evidence pointing out that the arrival times of interventions respond both to a set of time-varying covariates and the time elapsed since the last intervention. In particular, inflation changes and durations represent the strongest determinants of the timing of interest rate interventions. As regards asymmetric effects, we find that the Bank of Spain tends to intervene more frequently when: (i) inflation increases (ii) the exchange rate pta/DM depreciates, and (iii) the economy is undergoing a recessionary phase of the business cycle. Finally, we also find evidence in favor of the so-called 'duration effect', in the sense that the longer the period in which there is no intervention is, the larger the next interest rate change is.

The rest of the paper is organized as follows. Section 2 offers a brief explanation of the empirical methodology which is used throughout the paper and discusses the statistical framework under which the stochastic processes for *events* (*points)* and *marks* can be estimated. Section 3 analyzes the arrival times of central bank interventions through a *probit* model and tests for the possibility of asymmetries in its behavior. Section 4, in turn, analyzes interest rate changes through an *ordered probit* model. Section 5 tries to ascertain, though a *sequential probit* model, how robust are the results obtained in the previous sections to a joint estimation procedure. Finally, Section 6 concludes.

2 Econometric Issues

To measure the stance of monetary policy in Spain we use changes in the marginal interest rate of intervention of the Bank of Spain in the daily interbank market from 1984 to 1998. The choice of this sample period is dictated by the evidence saying that the Bank of Spain was controlling a broad monetary aggregate (ALP) as its main instrument of monetary policy before 1984 (see Escrivá and Santos, 1991 and Ayuso and Escrivá, 1996). The interest rate series is taken to be by the marginal interest rate of auctions "Prestamos de Regulación Monetaria" series from January 1984 to May 1990 and the interest rate of "Certificados de Depósito del Banco de España" (CEBES) series, from May 1990 to December 1998[1]. The evolution of that short-term interest rate is characterized by the following key features: (i) changes are spaced irregularly in time, (ii) target changes are in discrete increments. In the statistical literature, a stochastic process with such characteristics that has associated an information process at each event in time is known as a *marked point process*. The main goal of this section is to describe such processes, proceeding in two stages. At the first stage, we consider the arrival times of interventions (*events*) while at the second stage we analyse the size of changes in the interest rate (*marks*), given the decision to intervene.

A key issue in this analysis is how the timing of decisions evolves. In this respect, it is assumed that, at the beginning of each period, the central bank chooses a target for the interest rate, conditional on the information available up to that point, which is subsequently revised to accommodate the arrival of new information on the evolution on the economy. Under this scheme, given that the majority of macroeconomic data have monthly frequency, we will consider a month to be the relevant unit of time so that interventions will be aggregated to that fixed interval of time in the sequel.

[1] For a detailed description of the evolution of monetary policy in Spain during the sample period considered in this paper, see Servicio de Estudios del Banco de España(1997)

Thus, in this framework, it is convenient to begin by describing the statistical process which governs the decision of the Bank of Spain to intervene or not to intervene, given the available information set. The data can be viewed as follows. Let X_t be a random variable, that is assumed to be equal to zero if no intervention takes place during month t ($X_t = 0$), and equal to one if there is an intervention ($X_t = 1$). Let Y_t be the magnitude of the target change in month t, given that $X_t = 1$. Finally, consider B_{t-1} to be a Kx1 vector of predetermined variables known at the decision time (such as output growth, inflation growth, deviations of a monetary aggregate growth rate with respect to a targeted rate, exchange rate changes, etc.). In this way, the joint density of X_t and Y_t conditional on information up to time t-1 is:

$$X_t,Y_t/Z_{t-1} \sim f(X_t,Y_t/X_{t-1},Y_{t-1},B_{t-1},\theta) \qquad (1)$$

which without loss of generality, can be rewritten as the following product of marginal and conditional densities:

$$\begin{aligned} f(X_t,Y_t/X_{t-1},Y_{t-1},B_{t-1},\theta) &= f_1(X_t/X_{t-1},Y_{t-1},B_{t-1},\theta_1) * \\ & \quad f_2(Y_t/X_t,X_{t-1},Y_{t-1},B_{t-1},\theta_2) \end{aligned} \qquad (2)$$

That is, the joint density of *marks* and *events* is equal to the marginal density of the *events* (timing of central bank interventions), f_1, times the conditional density of the *marks*(changes in interest rates) given the current *event* , f_2, all conditional on past information. Under weak exogeneity in *marks*, then the MLE of θ_1 will be simply obtained from maximizing the marginal density in (2). Likewise, if X is weakly exogenous, the MLE of θ_2 will be obtained from maximizing the conditional density in (2).

Nevertheless, the possibility that weak exogeneity may not hold must be entailed. In this case, as Engle et al(1983) have argued, we will obtain consistent but inefficient estimates when only a single density is considered, for example, when we obtain estimates of θ_1 from f_1 and used them as given in f_2. However, this loss of efficiency tends to decrease as the sample sizes increases.

One additional problem is how to test for weak exogeneity of X for θ_2 in this econometric specification. If a *logit* specification is used to estimate both equations, then it is impossible to carry out an exogeneity test like, say, the one in Hausman (1978), since logistic sums are not logistic. For this reason, a possible solution is to choose a *probit* specification, since, due to the additive property of normality, this problem does not appear. However, even in such a case, a further problem arises when, as in this paper, the variable that creates the exogeneity problem, X, is a binary variable and therefore its distribution cannot be normal.

Due to those difficulties, we begin by reporting in sections 3 and 4 the results obtained through the independent estimation of a *probit* and an *ordered probit* for the marginal density in *events* and the conditional density of *marks*, respectively. Accordingly, the estimated coefficients reported below will be consistent but not neccesarialy efficient. Later, in Section 5, a *sequential probit* model is considered where the parameters of both densities are jointly estimated with the aim of ascertaining how robust are the results obtained with the first procedure.

3 Analysis of events : The arrival times of interventions

This section explains how to model the marginal density, denoted as f_1 in equation (2). The main goal is to explain how the arrival times of central bank ´s interventions are determined. In doing so, we aim to answer the following questions: *(i) What is the probability that the Bank of Spain will change the interest rate during the next month? (ii) Is there persistence in these interventions? (iii) What macroeconomic variables affect the timing of changes in the interest rate?*

To tackle those issues we use a *probit* model where it is assumed that interest rate changes take place at time 1,...,t-1, t , t+1,...where t refers to month t and T_t is the number of days elapsed from the last intervention. As aforesaid, the variable X_t equals one if the interest rate is changed in month t and equals zero otherwise. In this way, the probability of intervening can be defined in terms both of the number of days elapsed from the last intervention, and a set of explanatory variables. This can be written as:

$$p(X_t = 1/B_{t-1}, T_t, T_{t-1}) = F(\delta' H_t) = \Phi(\delta' H_t) \tag{3}$$

where $H_t = (B_{t-1}, T_t, T_{t-1})$, δ is a parameters vector, B_{t-1} are predetermined variables in monetary policy instrumentation and $\Phi(.)$ is the cumulative N(0,1) distribution[2].

[2] See Maddala (1983) for further details on the various probit estimated throughout this paper.

3.1 Empirical analysis

The main features of the data are briefly reported in Figures 1 to 3. Figure 1 plots the intervention intervals in days where the duration of no interventions appears in the vertical axis; for convenience we split the Figure into two panels: the first panel corresponds to the period 1984 to 1986, where the duration is short and there is a big outlier(see below) and the second panel refers to the period 1987 to 1998. Figure 2 reports the number of interventions by month. Lastly, Figure 3 presents the histogram of the daily interventions.

During the sample period there have been 700 interventions. The average duration between interventions is 7.85 days and duration almost doubles for the last part of the sample. The longest duration is 181 days and corresponds to the period between November 14, 1985 and March 5, 1986 when the transition from a monetary aggregate-based control scheme to an interest rate rule was being completed.

Next, we describe the choice of predetermined variables (B_{t-1}). Given the limited availability of macroeconomic data at monthly frequency, the following variables are considered: changes in the CPI inflation rate(π), changes in (log of)exchange rate pta/DM ($tcdm$), changes in the intervention interest rate(tm), deviations of a monetary aggregate (ALP) growth rate with respect to the targeted rate ($alp - alpo$) and changes in (log of) the Industrial Production Index (ipi)[3]. All variables are seasonally adjusted.

In principle all the predetermined variables are expressed in their absolute value. The main reason for using this transformation is that an initial

[3] Changes in registered unemployment were also considered without being significant in any of the estimated equations.

interesting exercise is to determine how the predetermined variables influence events in 'magnitude ', and not in 'sign '. Later, in section 3.2 the analysis will be extended to allow for asymmetries in the behaviour of the central bank.

Hence the initial specification of (3) becomes:

$$
\begin{aligned}
F(\delta' H_t) &= \Phi(\delta' H_t) = \Phi(\delta_0 + \delta_{d0} T_t + \delta_{d1} T_{t-1} + \delta_a |(alp - alpo)_{t-1}| + \\
&\quad \delta_{tc} |\Delta tcdm_{t-1}| + \delta_\pi |\Delta \pi_{t-1}| + \delta_p |\Delta ipi_{t-1}| + \delta_i |\Delta tm_{t-1}|)
\end{aligned} \tag{4}
$$

Table 1 reports the results obtained from estimating (4). Two probit models are in fact estimated: (i) one where responses are assumed to be the same over the whole sample period, and (ii) another which allows for different responses across relevant subperiods. For that, we intersect the regressors with a set of break-point dummy variables as follows: for $(alp - alpo)$ and T_t, we choose a step dummy variable in 1987:1 (d87) whereas for $\Delta tcdm$ and $\Delta \pi$ we choose a similar dummy variable with a break-point in 1992:01 (d92)[4]. The choice of those break-points is dictated by the following considerations. As mentioned above, the available evidence indicates that starting from the mid-eighties there was a change in the strategy of monetary policy control in Spain which culminated in 1987. Hence, that justifies the choice of the first date[5]. As for the second date, we introduce a break-point in 1992, since the Bank of Spain´s adoption of an inflation target and the changes in the pta/Ecu band fluctuation within the EMS took place in that year.

[4]For a detailed description of this methodology see Allison (1978) and Thompson (1977).

[5]See Ayuso and Escrivá(1996) and Escrivá and Santos(1991)

From the results in Table 1, it can be observed, as one would expect a priori, that the intervention probability responds positively and significantly (at least at 10% significance level) to changes in all the explanatory variables. Further it also depends positively on the time elapsed from the last intervention. The models fit appropiately, being the goodness of fit measures 0.50 and 0.61 in the model without and with breaks, respectively.

To ascertain the size of the previous effects in terms of altering the probability of intervention, the following comparison is made. First, assuming that there is no change in the set of explanatory variables, the intervention probability is estimated to be 0.22. Next, a change of 1 percentage point in each of the predetermined variables is separately considered. Table 2 summarizes the results of the simulations in both models. As regards the first model, the intervention probability increases from 0.22 to 0.32(inflation), 0.23(exchange rate), 0.24(money growth) and 0.23(industrial production), respectively. Further, to estimate the size of the duration effect, we have assumed that 50 days have elapsed since the last intervention, as compared to none. In this case the intervention probability increases from 0.22 to 0.33. Thus, both inflation changes and the duration seem to be the most important factors in explaining the arrival times of intervention.

With regard to the model which allows for regime breaks, we find that the intervention probability is 0.25 when there is no change in the set of explanatory variables. The inflation response is much larger after 1992 than before (0.30 and 0.42, respectively) in agreement with the Maastricht treaty requirement effects. By contrast, regarding the exchange rate pta/DM we find the opposite effect, namely, that this response is larger before 1992 than afterwards (0.32 and 0.25, respectively), given the increase in the width of the EMS currency fluctuation-band from 5 to 15 percentage points following the currency crisis in that year. As regards the effect of the time elapsed since the last intervention, we find that is marginally larger before 1987 than after 1987 (0.39 and 0.35, respectively). Finally, the responses to the lagged interest rate and monetary aggregate annual growth are found to be opposite, given the change that took place in the monetary policy instrument during the mid-1980s. Thus, for the interest rate, the probability is larger after 1987 than before (0.33 and 0.41, respectively) and the opposite happens with the monetary aggregate annual growth deviation from its target level (0.28 and 0.26, respectively).

3.2 Analysis of asymmetries

In this section, the previous analysis is extended to examine the possible existence of asymmetries in the decision to intervene by the Bank of Spain in response to changes in the set of explanatory variables. For that purpose, the *probit* model is respecified so that the coefficients on the predetermined variables are now allowed to depend on the *sign* of the changes. To do so, positive and negative changes are associated to two coefficients on each variable, δ^+ and δ^-, respectively. The null hypothesis of symmetry, $H_0 : \delta^+ = \delta^-$, is then tested using a Likelihood Ratio(LR) test. The results are as follows[6].

First, the probability of intervening is found to be larger when inflation accelerates than when it decelerates(p-value=0.01). This is to be expected

[6] Detailed results are available upon request.

since, during the sample period, both the reduction in the number of interventions and the disinflationary process took place simultaneously. Secondly, we observe a larger probability of intervention in the case of depreciations than of appreciations of the exchange rate(p-value=0.005). This indicates that when the peseta was getting appreciated, the Bank of Spain did not intervene so strongly as when it was depreciating. Thirdly, in response to changes in the industrial production index, the Bank of Spain seems to have exerted a stronger control during recessions than during expansions(p-value=0.04)[7]. Fourthly, no asymmetric behaviour is found in response to the deviations of a monetary aggregate growth rate with respect to its goal(p-value=0.08). This result could again be explained by the loss of significance of this variable in the monetary control strategy with the passage of time. Finally, it is found that interventions are less likely in situations when the interest rate increased in the last period than when it decreased(p-value=0.04). One possible interpretation of this result would be that the Bank of Spain gives less weight to the consequences on economic activity of a continuous rise in interest rate during an expansion than to a continuous decrease in a recession.

4 Analysis of marks: Size of interest rate changes

In this Section we model the density of the *marks*, denoted as f_2 in equation (2), conditional on the *events* and on the set of predetermined determinants of monetary policy instrumentation. Following Jordá(1997), we use an *ordered*

[7] Dolado and María-Dolores(1999) find evidence in favor of asymmetries effects on real activity depending on the cyclical state of the economy.

probit model since it can be viewed as a generalization of the linear regression model to cases where the dependent variable is discrete. In this way, it is only necessary to consider that the interest rate takes a finite number of values which possess a natural ordering.

For that, we assume that the *marks* of a *marked point processes* are only observed at each event time. That is, at times 1,...,i_t,...N we observe the magnitude of the target changes given by $Y_1, Y_2, ..., Y_N$. At any other time, Y is not observed. Let Y_i^* be an unobservable continuous random variable such that:

$$Y_i^* = w_i{}'\gamma + \varepsilon_i \qquad E[\varepsilon_i/w_i] = 0 \tag{5}$$

where $w_i = (T_i, Y_{i-1}, B_i)$ and ε_i is an error-term distributed as N(0,1). As before, the vector of variables B is assumed to be observed with one period lag. The assumption underlying the *ordered logit* model is that the observed target changes, Y_i, are related to the continuous latent variable, Y_i^*, in the following manner:

$$Y_i = \left[\begin{array}{ll} s_1 & \text{if } Y_i^* \in A_1 \\ s_2 & \text{if } Y_i^* \in A_2 \\ ... & \\ s_{m-1} & \text{if } Y_i^* \in A_{m-1} \\ s_m & \text{if } Y_i^* \in A_m \end{array} \right. \tag{6}$$

where the sets A_j partition the domain of Y_i^* such that $A_i \frown A_j = \emptyset$ for $i \neq j$.

The s_j are the observed discrete values of the variable Y_i and the partition of the domain of Y_i^* into subsets is such that:

$$\begin{aligned} A_1 &\equiv (-\infty, c_1) \\ A_2 &\equiv (c_1, c_2) \\ &.... \\ A_m &\equiv (c_{m-1}, \infty) \end{aligned} \tag{7}$$

Therefore, the distribution of the observed interest rate changes conditional on w_i becomes:

$$\begin{aligned} P(Y_i &= s_j \ / \ w_i) = P(w_i{}'\gamma + \varepsilon_i \ \in A_j) \\ &= \left[\begin{array}{ll} P(w_i{}'\gamma + \varepsilon_i \ \leq c_1) & if \ j = 1 \\ P(c_{j-1} < w_i{}'\gamma + \varepsilon_i \leq c_j \ ) & if \ 1 < j < m \\ P(c_j < w_i{}'\gamma + \varepsilon_i \ ) & if \ j = m \end{array} \right. \end{aligned} \tag{8}$$

which, under the assumption of normality, yields:

$$F(\eta) = \Phi(\eta) = \int_{-\infty}^{\eta} \exp(-\varepsilon^2/2) d\varepsilon \text{ for } \eta = c_j - w_i{}'\gamma \tag{9}$$

and then:

$$\left[\begin{array}{ll} \Phi(\varepsilon_i \ \leq c_1 - w_i{}'\gamma) & if \ j = 1 \\ \Phi(\varepsilon_i \leq c_j - w_i{}'\gamma \ ) - \Phi(\varepsilon_i \leq c_{j-1} - w_i{}'\gamma \ ) & if \ 1 < j < m \\ 1 \text{ - } \Phi(\varepsilon_i \leq c_{j-1} - w_i{}'\gamma \ ) & if \ j = m \end{array} \right. \tag{10}$$

The intuition behind the ordered logit model is clear: the probability of any particular observed target change is determined by where the conditional mean lies relative to the partition boundaries, c_j. Given this partition, a higher conditional mean implies a higher probability of observing a more extreme positive state. Hence, for example, if a large increase in inflation leads to a higher interest rate, we should expect the coefficients associated with the inflation variable to be positive.

To specify the likelihood, an indicator variable, I_i^j, which takes the value of 1 if $Y_i = s_j$ and 0 otherwise, is used. Then, the log-likelihood of the changes, conditional on the explanatory variables becomes[8]:

$$\pounds(Y/w) = \sum_{i=1}^{N} \left\{ \begin{array}{c} I_i^1 \log \Phi(c_1 - w_i{}'\gamma\ /\sigma) + \\ \sum_{j=1}^{m-1} I_i^j \log[\Phi(c_k - w_i{}'\gamma\ /\sigma) - \Phi(c_{k-1} - w_i{}'\gamma\ /\sigma)] \\ +I_i^m \log[1 - \Phi(c_{m-1} - w_i{}'\gamma\ /\sigma)] \end{array} \right\} \quad (11)$$

4.1 Empirical results

An *ordered probit* model allows to determine the ´magnitude ´ and the ´direction ´ of target changes. The estimated model is conditioned on the event taking place at time i_t and on information about the state of the economy at that time. In order to implement the model, the first step is to reduce the number of categories of the observed interest rate changes. Since in the data the average change is around 25 basis points, we have used the following classification:

[8] σ is not separately identifiable. Then, the estimation proceeds under the standard normalization $\sigma = 1$.

$$\begin{cases} \text{if } y_i \leq -0.625 \text{ then } \tilde{y_i} = -0.75 \\ \text{if -0.625} < y_i \leq -0.375 \text{ then } \tilde{y_i} = -0.5 \\ \text{if -0.375} < y_i \leq -0.125 \text{ then } \tilde{y_i} = -0.25 \\ \text{if -0.125} < y_i \leq 0.125 \text{ then } \tilde{y_i} = 0 \\ \text{if } 0.125 < y_i \leq 0.375 \text{ then } \tilde{y_i} = 0.25 \\ \text{if } 0.375 < y_i \leq 0.625 \text{ then } \tilde{y_i} = 0.5 \\ \text{if } y_i > 0.625 \text{ then } \tilde{y_i} = 0.75 \end{cases}$$

where $\tilde{y_i}$ refers to the aggregated series while y_i is the original series. As a result of this procedure, six intercepts need to be estimated. Thus we have:

$$y_i^* = w_i`\gamma + \varepsilon_i \qquad where\ \varepsilon_i \sim nid(0, \sigma)$$

where $\tilde{y_i} = \{-0.75, -0.5, -0.25, 0, 0.25, 0.5, 0.75\}$. Given the composition of B_i in the previous section, we specify (5) as follows:

$$\begin{aligned} y_i^* &= \gamma_y\, \tilde{y}_{i_{t-1}} + \gamma_T T_{i_t} + \gamma_a (alp - alpo)_{i_{t-1}} + \gamma_{tc} \Delta tcdm_{i_{t-1}} \\ &\quad + \gamma_\pi \pi_{i_{t-1}} + \gamma_{ipi} ipi_{i_{t-1}} + \varepsilon_i \end{aligned} \tag{12}$$

The estimation results are presented in Table 3 where it can be observed that the signs of the estimated coefficients associated to the different determinants are broadly in agreement with a priori beliefs. As before, we allowed for break-points in 1987 and 1992. We find that changes in inflation rate, the industrial production index, deviations of a monetary aggregate growth rate with respect to a target rate and a strong depreciation of exchange rate pta/DM tend to increase the size of interest rates changes. The estimated effect of inflation is larger after 1992, whereas for the remaining variables we find them to be smaller after 1987(money growth and duration) and 1992(exchange rates). As for the lagged change in the interest rate, we find its coefficient positive and strongly significant[9]. Finally, the coefficient associated with the duration variable, γ_T, shows that the higher duration is, the larger the change in the interest rate is[10].

In this respect, as Jordá(1997) points out, an important hypothesis to test is the so-called ´duration effect ´ that suggests that long spells without intervention might cause larger changes in interest rates than frequent adjustments. For this reason we construct the following indicator variable discriminating between shorter and larger duration than 15 days:[11]

$$\begin{cases} d_i^l = 1 \text{ if } dur > 15 \\ d_i^l = 0 \text{ otherwise} \end{cases} \textit{and} \begin{cases} d_i^s = 1 \text{ if } dur <= 15 \\ d_i^s = 0 \text{ otherwise} \end{cases} \tag{13}$$

Next, the following variables are defined: $T_i^l = d_i^l * T_i$ and $T_i^s = d_i^s * T_i$. Let γ_T^l and γ_T^s be coefficients associated with T_i^l and T_i^s. The hypothesis γ_T^l $= \gamma_T^s$ it is tested to check the existence of such an effect. Estimating the model and using a LR test we find evidence in favor of the duration effect, with the null hypothesis being rejected with a p-value of 0.007.

[9] This is in accordance with the obtained result by Rudebusch(1995) for the Fed-Funds rate

[10] The goodness of fit in this model is 0.441 but when we aggregate data in the following three categories(negative, zero and positive) it increases to 0.6466.

[11] The choice of 15 days corresponds to about two standard deviations of the duration data.

Finally, there is the possibility of testing for yet another type of asymmetry, namely, whether the ´duration´ effect depends on whether the interest rate is increasing or decreasing. Testing for this effect might be interesting since it is sometimes argued that being late when cuts in interest rates are needed may be more costly, in terms of achieving economic stability, than being late in rising interest rate. In other words it might be worse not to provide a stimulus to the economy when it is in a recession than not to delay putting a brake to an overheated economy. For that, we define the following indicator variable that depends on whether the industrial production index growth rate is positive or negative:

$$\begin{cases} d_i^+ = 1 \text{ if } \Delta ipi_i \geq 0 \\ d_i^+ = 0 \text{ otherwise} \end{cases} and \begin{cases} d_i^- = 1 \text{ if } \Delta ipi_i < 0 \\ d_i^- = 0 \text{ otherwise} \end{cases} \tag{14}$$

In the same way as we did before we define: $T_i^+ = d_i^+ * T_i$ and $T_i^- = d_i^- * T_i$ where γ_T^+ and γ_T^- are the coefficients associated with T_i^+ and T_i^-. When the null hypothesis $\gamma_T^+ = \gamma_T^-$ is tested, we find that it is again rejected with a p-value of 0.02. In fact, though not reported, the estimated coefficients on T_i^- are insignificant while the coefficients on $T_i^+$are strongly significant, a result which would seem to suggest that, over the sample period, the Bank of Spain has been more keen in avoiding overheating of the Spanish economy than in fighting recessions.

5 Events and Marks: A sequential probit model

In the previous sections, both the marginal and the conditional distributions, $f_1$and f_2 in (2), have been estimated separately. To test wether the obtained results remain robust to estimating them jointly, we use a *sequential probit* model by means of which we estimate jointly the parameters in the densities for *events* and *marks*.

We assume that the central bank makes a joint decision on whether to intervene ($X = 1$) or not to intervene ($X = 0$) and, in the former case ,on the size of the intervention which, as before, is assumed to take the values $Y = -0.75, -0.5, -0.25, 0, 0.25, 0.5, 0.75$. Hence, assuming that each choice can be done on the basis of a probit model, the *sequential probit* can be specified in the following manner:

$$p(X_t = 0/B_{t-1,}T_t, T_{t-1}) = \Phi(-\delta H_t) \tag{15}$$

in the case of no intervention and

$$p(Y_i = s_j/w_i) = p(X_t = 1/B_{t-1,}T_t, T_{t-1}) * p(w_i`\gamma + \varepsilon_i \in A_j) \tag{16}$$

in the case of intervention. This leads to the following likelihood function, where j is defined as above.

$$\left[\begin{array}{ll} (1\text{-}\Phi(-\delta H_t))\Phi(\varepsilon_i \ \leq c_1 - w_i{}'\gamma) & if\ j = 1 \\ (1\text{-}\Phi(-\delta H_t))(\Phi(\varepsilon_i \leq c_j - w_i{}'\gamma\ ) - \Phi(\varepsilon_i \leq c_{j-1} - w_i{}'\gamma\ )) & if\ 1 < j < m \\ (1\text{-}\Phi(-\delta H_t))(1 \text{ - } \Phi(\varepsilon_i \leq c_{j-1} - w_i{}'\gamma\ )) & if\ j = m \end{array}\right. \tag{17}$$

Hence, the parameters of the model are estimated through maximization of the joint likelihood function in equations (15) and (17).

5.1 Empirical results

Table 4 shows the estimated parameters for the *sequential probit* model. Though a proper statistical test is not available, it can be noticed that the results are fairly similar to those obtained separately in the probit (for *events*) and ordered probit(for *marks*) models[12]. So, for example, and as before, we obtain a larger probability of intervention when inflation accelerates than when it decelerates(p-value= 0.0005), when the exchange rate pta/DM is depreciating than when it is appreciating(p-value=0.01), when there is excessive monetary growth (p-value= 0.02) and finally when the lagged interest rate is increasing (p-value= 0.01).

Lastly we tested once more for ´duration effects ´ in this model. Similar results are obtained: both effects seem to be present with p-values of 0.006 and 0.04, respectively. However, with regard to industrial output changes, we do not find in this case any clear asymmetries(p-value=0.06).

6 Conclusions

In this paper we have studied the determinants of the arrival times and the size of the Bank of Spain´s interventions, in terms of interest rate changes

[12]Moreover, the correlation between the residuals in the estimated models for *events* and marks is small(0.1345), which also supports the consistency of estimates obtained in the separate.

in the daily interbank market, through the estimation of a *marked point process*. In particular, at a first stage, the effect of various determinants of the timing of interventions by the central bank(*events*) are estimated while, at a second stage, we analyze the role of those variables in determining the size of interventions(*marks*). In both cases probit specifications are used.

Since a caveat of this methodology is that an exogeneity problem may appear when estimating both densities independently, we find it appropriate to use a *sequential probit* model to estimate the joint likelihood of *events* and *marks* in order to check for consistency of the estimates obtained in the separate estimation approach. Moreover, we test for asymmetries in the response of the Bank of Spain to positive and negative changes in the variables contained in its information set.

The results regarding the *events* can be summarized as follows. It is observed that there is a higher probability of intervention when: (i) there is a recession, (ii) the inflation rate is increasing, (iii) the exchange rate is depreciating, and (iv) the monetary aggregate growth is under its annual target rate. In particular, the response to inflation changes seems to be much larger after 1992 than before.

As for the *marks* we find evidence in favour of a ´duration effect ´. This implies that large inactivity periods induce stronger changes in interest rates than frequent adjustments. Finally, the obtained results point out than the Bank of Spain has been more prone to avoid overheating, by implementing large increases in interest rates during expansions, that to relax monetary policy stance during recessionary periods, through large cuts in interest rates.

Table 1		
A probit model for central bank interventions		
Sample period: January 1984 - December 1998		
Variable	**Coefficient**[1]	**Coefficient**[2]
T_t	0.011	0.013
	(1.53)	(2.29)
$T_t * d87$	-	-0.003
	-	(1.52)
T_{t-1}	0.021	0.088
	(3.24)	(2.48)
$T_{t-1} * d87$	-	-0.019
	-	(2.25)
$\|(alp - alpo)\ _{t-1}\|$	0.091	0.198
	(2.88)	(1.69)
$\|(alp - alpo)\ _{t-1}\| * d87$	-	-0.119
	-	(2.08)
$\|\Delta tcdm_{t-1}\|$	0.024	0.364
	(2.31)	(1.66)
$\|\Delta tcdm_{t-1}\| * d92$	-	-0.346
	-	(1.70)
$\|\Delta \pi_{t-1}\|$	0.482	0.282
	(2.46)	(1.94)
$\|\Delta \pi_{t-1}\| * d92$	-	0.527
	-	(2.66)
$\|\Delta ipi_{t-1}\|$	0.016	0.013
	(1.73)	(1.67)
$\|\Delta tm_{t-1}\|$	0.392	0.410
	(2.97)	(1.65)
$\|\Delta tm_{t-1}\| * d87$	-	0.354
	-	(2.45)
$constant$	-1.239	-1.118
	(12.17)	(9.44)
$Log - Likelihood$	337.97	358.30
No of observations	739.	

Note:(1) and (2) denotes a probit model without and with break-points, respectively. Heteroskedasticity-robust t-ratios in parentheses

Table 2			
Effects of changes in predetermined variables			
on intervention probability(ρ)			
Model without break-points			
Change	**Initial probability**	**Final probability**	
$T_t = 50$	0.225	0.334	
$\lvert (alp - alpo)\ _{t-1} \rvert = 0.01$	0.225	0.240	
$\lvert \Delta tcdm_{t-1} \rvert = 0.01$	0.225	0.228	
$\lvert \Delta \pi_{t-1} \rvert = 0.01$	0.225	0.319	
$\lvert \Delta ipi_{t-1} \rvert = 0.01$	0.225	0.227	
$\lvert \Delta tm_{t-1} \rvert = 0.01$	0.225	0.300	
Model with break-points(b-p)			
Change	**Initial ρ**	**Final ρ before b-p**	**Final ρ after b-p**
$T_t = 50$	0.246	0.385	0.350
$\lvert (alp - alpo)\ _{t-1} \rvert = 0.01$	0.246	0.284	0.261
$\lvert \Delta tcdm_{t-1} \rvert = 0.01$	0.246	0.319	0.249
$\lvert \Delta \pi_{t-1} \rvert = 0.01$	0.246	0.302	0.423
$\lvert \Delta ipi_{t-1} \rvert = 0.01$	0.246	0.249	0.249
$\lvert \Delta tm_{t-1} \rvert = 0.01$	0.246	0.330	0.412

Table 3

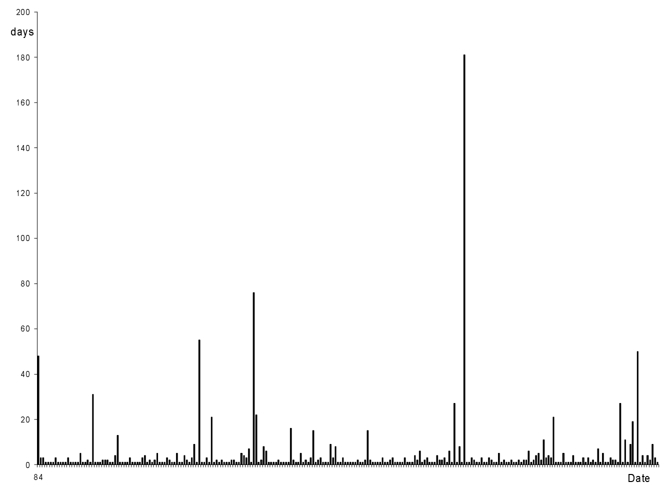
An ordered probit model for changes in intervention interest rate

Sample period: January 1984 - December 1998

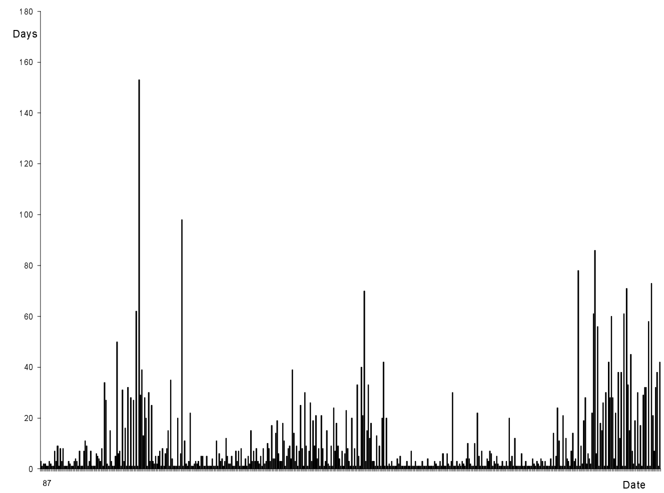
Coefficient	**Estimated Value**	**t-statistic**
γ_y	0.245	2.99
γ_T	0.023	1.69
$\gamma_T * d87$	-0.011	2.23
γ_a	0.272	2.38
$\gamma_a * d87$	-0.200	2.05
γ_{tc}	0.491	1.71
$\gamma_{tc} * d92$	-0.467	1.63
γ_π	0.051	2.50
$\gamma_\pi * d92$	0.041	1.92
γ_{ipi}	0.023	1.91
c_1	-1.275	6.70
c_2	-0.859	4.67
c_3	0.156	1.86
c_4	1.047	5.73
c_5	2.059	10.39
c_6	2.299	11.68
Log-Likelihood	1135.22	-
No.of observations	700	

Table 4			
A sequential probit model for central bank interventions			
Sample period: January 1984 - December 1998			
Coefficient	**Probit**	**Coefficient**	**Ordered Probit**
T_t	0.009 (1.71)	γ_y	0.262 (2.37)
$T_t * d87$	-0.0059 (1.76)	γ_T	0.013 (2.73)
T_{t-1}	0.005 (3.06)	$\gamma_T * d87$	-0.001 (1.94)
$T_{t-1} * d87$	0.0071 (2.23)	γ_a	0.502 (2.05)
$\|(alp - alpo)\ _{t-1}\|$	0.2176 (2.81)	$\gamma_a * d87$	-0.419 (1.66)
$\|(alp - alpo)\ _{t-1}\| * d87$	-0.0906 (1.79)	γ_{tc}	0.093 (2.08)
$\|\Delta tcdm_{t-1}\|$	0.4664 (2.49)	$\gamma_{tc} * d92$	-0.092 (2.64)
$\|\Delta tcdm_{t-1}\| * d92$	-0.4438 (1.78)	γ_π	0.067 (2.10)
$\|\Delta \pi_{t-1}\|$	0.8192 (2.63)	$\gamma_\pi * d92$	0.066 (3.94)
$\|\Delta \pi_{t-1}\| * d92$	0.0433 (1.87)	γ_{ipi}	0.003 (1.71)
$\|\Delta ipi_{t-1}\|$	0.016 (1.90)	c_1	-1.08 (17.99)
$\|\Delta tm_{t-1}\|$	0.4189 (2.95)	c_2	-0.65 (16.56)
$\|\Delta tm_{t-1}\| * d87$	0.3237 (2.33)	c_3	0.18 (7.51)
$constant$	-1.101 (13.82)	c_4	1.60 (11.48)
$Log - Likelihood$	1711.38	c_5	2.69 (17.44)
No. of observations	739	c_6	3.74 (13.10)

Note: Heteroskedasticity-robust t-ratios in parentheses



Sample Period:1984-1986



Sample Period:1987-1998

Figure 1: Duration in central bank intervention in interest rates(days)

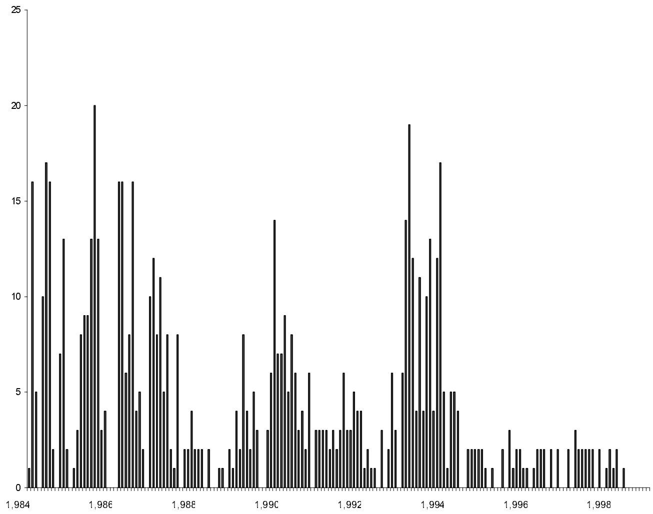


Figure 2: Number of interventions in interest rate by month

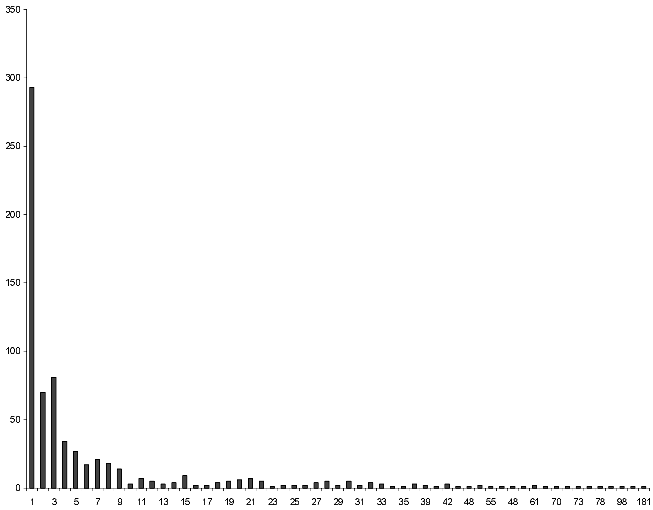


Figure 3: Duration in days histogram

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